Market Entries and Exits and the Nonlinear Behaviour of the Exchange Rate Pass-Through into Imported Prices

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Abstract

This paper develops an empirical framework giving rise to a nonlinear behaviour of the exchange rate pass-through (ERPT). Rather than shifts between low and high inflation, the nonlinearity arises when large swings in the exchange rate trigger market entries and exits of importing firms. Switching regressions are used to distinguish between low and high pass-through regimes of the exchange rate into import prices. For the case of Switzerland, the corresponding results suggest that, though inflation has been low and stable, the ERPT still doubles in value in times of a rapid appreciation of the Swiss Franc.

JEL classification: F15, F31, L11

Keywords: Exchange Rate Pass-Through, Import Prices, Switching Regression

1 Introduction

> The pass-through coefficient in the phase where the number of foreign firms is constant is quite small, perhaps close to zero, whilst its values in the phases with entry or exit are much larger, perhaps close to one. Dixit (1989, p.227)

In several areas of economic policy, the exchange rate pass-through (ERPT)—that is the impact of price changes of foreign currency upon import prices measured in terms of an elasticity—is a carefully watched variable. For example, for monetary policy, the value of the ERPT connects the developments on the foreign exchange market with inflation whilst, for antitrust policy, the ERPT indicates in how far import competition prevents local producers from charging excessively high prices. It is therefore not surprising that a plethora of empirical research has been devoted to estimating the pass-through effect. Campa and Goldberg (2005), Ihrig *et al.* (2006), and Frankel *et al.* (2011) provide some recent examples reporting estimates for several countries. Numerous other studies have dealt with the conditions of individual countries and industries. Probably the most important stylised fact arising from this research is that, even in the long-term, import prices adjust incompletely to exchange rates. Elasticities of around -0.5 are commonly found (Goldberg and Knetter, 1997).

Market frictions are essential to explain why import prices react partially to changes in currency prices. Such frictions include price rigidities, which are a key ingredient in open economy models of the ERPT (see Taylor, 2000; Choudhri and Hakura, 2006; or Devereux and Yetman, 2010). Though it has long been recognised that, with sticky prices, different levels of inflation can give rise to differences in the ERPT between countries (e.g. Taylor, 2000; Gagnon and Ihrig, 2004; Campa and Goldberg, 2005; Choudhri and Hakura, 2006; Frankel *et al.*, 2012), only recently, Al-Abri and Goodwin (2009) as well as Shintani *et*

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al. (2013) have suggested that this might warrant nonlinear models to estimate the nexus between import prices and exchange rates. Arguably, when menu cost create a threshold delineating when an adjustment of import prices is worthwhile, the ERPT of a given country could vary between periods with low and high inflation.

This paper endeavours to contribute to the empirical literature measuring the ERPT by considering that nonlinearities can also arise from changes in competition when importing firms start to enter or exit a given market. Imperfect competition has provided a second explanation for why foreign firms only partially adjust their import prices to changes in the exchange rate. Dornbusch (1987) has pioneered the literature embedding the ERPT into models of industrial organisation. A strand of the literature tying the incomplete ERPT with oligopolistic market structures and product differentiation has allowed for the possibility that market entries and exits by foreign firms alter the degree of import competition. Though they pursue slightly different theoretical approaches, the seminal contributions of Baldwin (1988) and Dixit (1989) both emphasise the combined role of irreversible entry and exit cost and exchange rate uncertainty in guiding the decision of foreign firms to, respectively, start or terminate supplying goods to a given market. As illustrated by the quote at the outset, a key result of this literature is that exchange rates can impact upon prices in a nonlinear manner. In particular, during periods with relatively stable exchange rates, even modest sunk cost could discourage foreign firms from changing the status quo. With a fixed market structure, competition is limited to the setting of prices and the ERPT might be relatively low. Conversely, sufficiently large swings in the foreign exchange market affect the profits of importers to a degree where they will want to incur the irreversible cost to enter or exit a market. The resulting adjustment in the share of imported goods gives rise to an additional channel forcing foreign firms to adapt their prices according to the conditions on the foreign exchange market. Though the role of openness to international trade has received attention in the empirical ERPT literature (McCarthy, 2000; Gust *et al.*, 2010; An and Wang, 2011), the nonlinearities arising in the theoretical work of Baldwin (1988) and Dixit (1989) from changes in the market structure have hitherto been ignored.

To fill this gap, this paper develops a stylised model where the ERPT has the widely found elasticity of around -0.5 when the market structure is fixed, but higher values when large shifts in the foreign exchange market induce foreign firms to enter or exit the market. Consistent with Baldwin (1988) and Dixit (1989), the theoretical framework gives rise to a low and high pass-through regime. Empirically, these regimes can be represented by switching regressions. The possibility of a regime change is illustrated for the case of Switzerland, which has been chosen due to its low and stable level of inflation during the last decades. Nevertheless, an unobserved regime switching regression suggests that the effect of the exchange rate is not constant but doubles in value in times of marked appreciations of the Swiss Franc, when inflation was in general low and, hence, a menu cost explanation is unlikely to apply.

The paper is organised as follows. Section 2 derives the model distinguishing different regimes of the ERPR depending on whether or not firms enter or exit the market. Section 3 connects this theoretical framework with a regime switching regression. Section 4 presents an empirical application with data from Switzerland. Section 5 summarises and concludes.

2 Market Entry and Exchange Rate Pass-Through

This section develops a stylised model where the market entry and exit of importing firms gives rise to nonlinear transmission effects of the exchange rate onto import prices. Rather than providing an in depth discussion about this topic, which has appeared in the seminal work on the hysteresis effects of the exchange rate (Baldwin, 1988; Dixit, 1989; Baldwin and Krugman, 1989), the aim is merely to develop a framework underpinning the empirical analysis of Sections 3 and 4.

Consider a model where the domestic market of a given country is served by a number of identical foreign importers, whose weight equals ω_t, whilst $(1-\omega_t)$ of the market is covered by identical local producers. Furthermore, during period t, the importing and local sector charge prices denoted by, respectively, p_t^* and p_t.

The demand conditions are modeled with the transcendental logarithmic (or translog) expenditure function. Translog-preferences have the distinctive property that the market share between, say, imported and locally produced goods is *not* constant, but depends on the differences between the corresponding prices p_t^* and p_t (see Bergin and Feenstra, 2000, 2001). This is maybe a crucial ingredient when considering the effect of a changing market structure, where importers can enter or exit the market. As derived in Herger (2012), for a scenario of representative local and importing firms, the expenditure E_t at time t with a translog function is given by

$$\begin{aligned}\ln E_t(\overline{U}, p_t^*, p_t) &= \omega_0 + \ln \overline{U} + \omega_t \ln p_t^* + (1-\omega_t)\ln p_t \\ &- \frac{1}{2}\gamma^* \ln p_t^* \ln p_t^* + \frac{1}{2}\gamma^* \ln p_t^* \ln p_t + \frac{1}{2}\gamma \ln p_t \ln p_t^* - \frac{1}{2}\gamma \ln p_t \ln p_t\end{aligned}$$

where ω_0 is a constant, $\overline{U}$ the utility level to be reached with expenditure E at prevailing prices p_t^* and p_t, and γ reflects the degree of substitutability between locally produced goods and imports (and vice versa for γ^*). Shephard's Lemma[1] implies that the expenditure share s_t^* on imports is given by

$$s_t^*(p_t^*, p_t) = \frac{\partial \ln E_t}{\partial \ln p_t^*} = \omega_t - \gamma^* \ln p_t^* + \underbrace{\frac{\gamma+\gamma^*}{2}}_{=\gamma^*\Gamma} \ln p_t. \tag{1}$$

Unlike in Bergin and Feenstra (2001), the symmetry assumption $\gamma^* = \gamma$ is here *not* introduced. Following Herger (2012, p.385), the degree of asymmetry between the (countervailing) effects of import and local producer prices on the share of imports is summarised by $\gamma^*\Gamma = (\gamma+\gamma^*)/2$ with $\Gamma = 1$ reflecting symmetric conditions whilst $\Gamma > 1$ and $0 < \Gamma < 1$ indicate, respectively, a relatively high and low sensitivity of s_t^* to local producer prices. As mentioned above, even with a constant weight ω_t of importing firms, the expenditure share s_t^* is not fixed but depends on the price difference between locally produced and imported goods. The price elasticity of imports depends on the degree of substitutability embodied in γ and γ^* with the limit value of 0 reflecting the CES-case with perfect complements.

For an importing firm, the price p_t^* of selling a product abroad and the production cost are denominated in different currencies. With the nominal exchange rate e_t—expressed here as the price between the currency of the importing firm and the domestic currency—the profit of an established importer equals

$$\pi_t^* = p_t^* - (1/e_t), \tag{2}$$

where, for the sake of simplicity, the quantity of imports and foreign production costs have been normalised to 1. Hence, an increase in e_t, which is an appreciation of the domestic currency, reduces the cost of foreign production relative to the import price p_t^*. From this, as derived in appendix A.1, the optimal price setting rule is given by

$$p_t^* = \left[1 + \frac{s_t^*}{(1-\nu\Gamma)\gamma^*}\right]\frac{1}{e_t}. \tag{3}$$

[1] According to Shephard's Lemma, the Hicksian demand function equals $h = \partial E_t/\partial p_t^*$ and the market share equals

$$s_t^* = \frac{h_t^* p_t^*}{E_t} = \frac{\partial E_t}{\partial p_t^*}\frac{p_t^*}{E_t} = \frac{\partial \ln E_t}{\partial \ln p_t^*}.$$

where $\nu = \frac{\partial p_t}{\partial p_t^*} \frac{p_t^*}{p_t}$ is the conjectural elasticity reflecting the percentage change of local prices that importers expect when they change their price by one per cent. As discussed in Bernhoven and Xu (2000), the conjectural elasticity provides a concise way to account for the impact of various degrees of price competition upon the ERPT. In particular, we have that $-\infty < \nu < 1$ with higher values reflecting less competition in the sense that an importing firm can push prices p_t^* further above the marginal cost $1/e_t$. The special case of Cournot-Nash behaviour implies that $\nu = 0$ by assumption. Taking logarithms of (3) and using the first order Taylor approximation yields $\ln p_t^* \approx \frac{s_t^*}{(1-\nu\Gamma)\gamma} - \ln e_t$. Inserting (1) for s_t^*, and solving for import prices, yields the optimal pricing equation

$$\ln p_t^* \approx \frac{1}{(2-\nu\Gamma)} \ln p_t - \frac{1-\nu\Gamma}{2-\nu\Gamma} \ln e_t - \frac{1}{(2-\nu\Gamma)\gamma^*}\omega_t. \tag{4}$$

Reflecting the seminal result of Dornbusch (1987), in (4), the ERPT given by $(1-\nu\Gamma)/(2-\nu\Gamma)$ depends on degree of price competition. In particular, perfect competition with $\nu = -\infty$ implies that $(1-\nu\Gamma)/(2-\nu\Gamma) = 1$ and the adjustment of import prices to exchange rate movements is instantaneous and complete. Conversely, under a scenario with perfectly matched price setting across local and foreign firms with $\nu = 1$ and $\Gamma = 1$, there is no effect of the exchange rate upon import prices. Note that Cournot-Nash behaviour with $\nu = 0$ and symmetric reactions with $\Gamma = 1$ give rise to a pass-through effect of $1/2$. Implicitly, this parametrisation appears in Bergin and Feenstra (2000, pp.662-663) and is also close to the values that are widely found in empirical work (Goldberg and Knetter, 1997).

Competition might not only depend on the extend with which importing firms react to the price setting of existing rivals, but also on the prospect that profits and losses can trigger market entries and exits. The pass-through effect discussed in the previous paragraph has neglected the effect of the exchange rate on the weight ω_t of importing firms. However, as discussed at the outset, it has been argued that market entries and exits occur predominantly in times of large swings in the foreign exchange market, since episodes of marked depreciations or appreciations can have profound effects on the earnings of foreign firms. Inserting the optimal price of (3) into (2) implies that the profit function equals $\widetilde{\pi}_t^* = (s_t^*/(\gamma^*(1-\nu\Gamma))(1/e_t)$. As shown in appendix A.2, inserting (1) for s_t^* yields approximately

$$\ln \widetilde{\pi}_t^* \approx \frac{1}{(1-\nu\Gamma)\gamma^*}\omega_t - 1 - \frac{1}{1-\nu\Gamma} \ln p_t^* + \frac{\Gamma}{1-\nu\Gamma} \ln p_t. - \ln e_t \tag{5}$$

Of note, under the Cournot-Nash $\nu = 0$ and symmetry assumption $\Gamma = 1$, (5) implies that prices and exchange rates are tied together in a direct proportional relationship that is similar to the purchasing power parity condition (PPP). Furthermore, zero profits with $\ln \widetilde{\pi}_t^* = 1$ define a scenario where no foreign firm enters or exits the market, which gives rise to an equilibrium condition for the exchange rate given by $u_t = \frac{1}{(1-\nu\Gamma)\gamma^*}\omega_t - 2 - \frac{1}{1-\nu\Gamma} \ln p_t^* + \frac{\Gamma}{(1-\nu\Gamma)} \ln p_t - \ln e_t = 0$. Conversely, positive profits $\ln \widetilde{\pi}_t^* > 1$ arise when $u_t > 0$ indicating an overvaluation of the domestic currency in the sense that value of e_t is relatively low. Likewise, losses $\ln \widetilde{\pi}_t^* < 1$ indicate that the domestic currency is undervalued with $u_t < 0$ in the sense that value of e_t is relatively high.

Adjustments in the weight ω_t of importing firms occur when the profit and losses of (5) outweight the cost accruing, respectively, to entering and exiting the local market. Assume, for the sake of simplicity, that the logarithmically transformed cost of an importing firm that wants to switch its status between serving and exiting a market equals f. As long as $|\widetilde{\pi}_t^*| = |u_t| \leq f$ no firms enter or exit and ω_t is constant. Yet, sufficiently high profits or losses give rise to an adjustment of the market structure until $|\widetilde{\pi}_t^*| = |u_t| = f$. Solving (5) for the corresponding import share yields $\omega_t = \gamma^* \ln p_t^* - \gamma^*\Gamma \ln p_t + (1-\nu\Gamma)\gamma^* \ln e_t + c$ where c is a constant. Taken together, depending on $|\widetilde{\pi}_t^*| = |u_t| \lesseqgtr f$, we have that

$$\omega_t = \begin{cases} \omega_t & \text{if} \quad |u_t| \leq f \quad \text{No Entries/Exits} \\ \gamma^* \ln p_t^* - \gamma^* \Gamma \ln p_t + (1-\nu\Gamma)\gamma^* \ln e_t + c & \text{if} \quad |u_t| > f \quad \text{Enties/Exits.} \end{cases} \tag{6}$$

Since the import share ω_t adjusts as function of e_t, the ERPT exhibits a nonlinear behaviour around the points where entries and exits occur. In particular, as derived in appendix A.3, substituting (6) into (4) and rearranging yields

$$\Delta p_t^* \approx \begin{cases} [1/(2-\nu\Gamma)] \ln p_t & -\big[(1-\nu\Gamma)/(2-\nu\Gamma)\big] \ln e_t \quad -\big[1/((2-\nu\Gamma)\gamma^*)\big]\omega_t \quad \text{if } |u_t| \leq f \\ \big[(2\Gamma)/(3-\nu\Gamma)\big] \ln p_t & -\big[(2-2\nu\Gamma)/(3-\nu\Gamma)\big] \ln e_t \quad \text{if } |u_t| > f. \end{cases}$$

Subtracting lagged terms from both sides yields transforms this into log differences, with $\Delta p_t^* = \ln p_t^* - \ln p_{t-1}^*$, $\Delta p_t = \ln p_t - \ln p_{t-1}$, and $\Delta e_t = \ln e_t - \ln e_{t-1}$, that is

$$\Delta p_t^* \approx \begin{cases} \big[1/(2-\nu\Gamma)\big]\Delta p_t & -[(1-\nu\Gamma)/(2-\nu\Gamma)]\Delta e_t \quad \text{if } |u_t| \leq f \\ \big[(2\Gamma)/(3-\nu\Gamma)\big]\Delta p_t & -[(2-2\nu\Gamma)/(3-\nu\Gamma)]\Delta e_t \quad \text{if } |u_t| > f. \end{cases} \tag{7}$$

The first line of (7) refers to a scenario where prices and exchange rates are close to their equilibrium and hence only price competition from established rivals forces foreign firms to adjust their import prices to exchange rates (note that $\Delta\omega_t = 0$). This reflects the scenario discussed above after (4). However, sufficiently large swings in the foreign exchange rate result in profits and losses that end up changing the market structure. Then, the weight ω_t of the importing sector becomes an endogenous variable. The second line of (7) accounts for this channel, which fosters the ERPT since a sufficiently large over- or undervaluation of a currency can give rise to market entries and exits which, in turn, increase the competitive pressure to adjust the import prices to changes in the exchange rate. It is maybe again instructive to consider the case of Cournot-Nash behaviour with $\nu = 0$ and symmetry $\Gamma = 1$ which, in the second line of (7), gives rise to a pass-through elasticity of $2/3$.

The merit of (7) is to tie with Baldwin (1988) and Dixit (1989) who have emphasised the role of nonlinearities in the relationship between exchange rates and import prices. In sum, there might be two regimes in the ERPT with relatively larger effects arising in times of pronounced shifts in the value of the exchange rate. This possibility has been neglected in the empirical literature on the ERPT. The following sections will turn to this issue.

3 Econometric Strategy

This section develops an econometric framework that encapsulates the nonlinear effects of the ERPT discussed in Section 2. In particular, the switching regression[2] lends itself to tracing the distinction between a low-pass regime where price competition and the high-pass regime where also market entries and exits force foreign firms to adjust their prices according to the conditions on the foreign exchange market. The econometric equation reflecting the structure of (7) is given by

$$\Delta p_t^* = \beta_m^e \Delta e_t^* + \beta_m^p \Delta p_t + \epsilon_t(\sigma_m), \tag{8}$$

where β_m^e and β_m^p are coefficients to be estimated. The variables enter in logarithmic differences that is $\Delta p_t^* = \ln p_t^* - \ln p_{t-1}^*$, $\Delta e_t = \ln e_t - \ln e_{t-1}$, and $\Delta p_t = \ln p_t - \ln p_{t-1}$. The main difference between a standard regression equation and (8) is that the value of the coefficients can differ across regimes m where $m = 0$ denotes the low-pass and $m = 1$ the high-pass regime. Hence, the theoretical expectation is that $\beta_0^e < \beta_1^e$. Finally, $\epsilon_t(\sigma_m)$ is a stochastic error term whose standard deviation σ_m can be regime dependent.

[2]For a textbook discussion on regime switching regressions see Hamilton (1994, ch.22). Without looking at the role of different forms of import competition, Hernandez and Leblebicioğlu (2012) employ a Markov switching regression to capture the nonlinear reactions in the ERPT for cars imported into the US market.

In practice, the postulated ERPT-regimes $m \in 0, 1$ are not directly observable. However, the theoretical framework of the previous section resulting in (7) indicates that the probability $P[m_t = i]$ of observing regime $i = 0, 1$ during period t depends on the degree of over- or undervaluation u_{t-1} of a currency. Regime switching models differ with respect to the definition of this probability. In a simple case, the probabilities $P[m_t = i]$ are independent of each other and depend on u_{t-1} according to a multinomial logit distribution, that is

$$P_i[m_t = i|u_{t-1}, \delta_i] = \frac{\exp(\delta_i u_{t-1})}{1 + \exp(\delta_i u_{t-1})} \tag{9}$$

with δ_i denoting a regime specific coefficient with normalisation $\delta_0 = 0$. The special case where $u_{t-1} = 1 \;\forall\; t$ yields a constant regime probability. A more popular specification assumes that regimes follow each other according to a Markov-chain with $P(m_t = i|m_{t-1} = j)$ with $j = 0, 1$. Under this scenario, a matrix of transition probabilities arises that contains all four contingencies of moving between a low and high-pass regime ($P[m_t = 0|m_{t-1} = 0], P[m_t = 0|m_{t-1} = 1], P[m_t = 1|m_{t-1} = 0], P[m_t = 1|m_{t-1} = 1]$). Similar to the discussion above, the transition probability can depend on observable variables according to a multinomial logit distribution, that is

$$P_i(m_t = i|m_{t-1} = j, u_{t-1}, \delta_{ij}) = \frac{\exp(\delta_{ij} u_{t-1})}{1 + \exp(\delta_{ij} u_{t-1})} \tag{10}$$

which involves four coefficients δ_{ij} as well as the normalisation $\delta_{01} = \delta_{11} = 0$.

Assuming that the stochastic error term follows a standard normal distribution, that is $\epsilon_t \sim N(0, \sigma_m)$, the likelihood contribution of an observation at time t equals $L_t = \sigma_m^{-1}(\Delta p_t^* - \beta_0^e \Delta e_t^* - \beta_0^p \Delta p_t)P[m_t = 0] + \sigma_m^{-1}(\Delta p_t^* - \beta_1^e \Delta e_t^* - \beta_1^p \Delta p_t)P[m_t = 1]$. The log-likelihood function, from which the coefficients β can be estimated, equals then

$$\ln L = \sum_{t=1}^{T} \ln \left\{ \phi\left(\frac{\Delta p_t^* - \beta_0^e \Delta e_t^* - \beta_0^p \Delta p_t}{\sigma_0}\right) P[m_t = 0] + \phi\left(\frac{\Delta p_t^* - \beta_1^e \Delta e_t^* - \beta_1^p \Delta p_t}{\sigma_1}\right) P[m_t = 1] \right\} \tag{11}$$

with ϕ denoting the probability density function of the standard normal distribution. An exact specification of (11) arises when (9) or (10) define the probability $P[m_t = i]$. However, (11) depends on the one-step ahead probability $P[m_t = i|u_{t-1}]$, wherefore the maximisation occurs via a recursive evaluation.[3]

Finally, since u_{t-1} is not directly observable, it needs estimating from the data. It is well known from the cointegration literature that equilibrium relationships that tie variables together, such as the one postulated by (5), can be uncovered by regressing the *level* of e.g. the import prices $\ln p_t^*$ onto the *level* of the exchange rate $\ln e_t$ and the domestic producer prices $\ln p_t$, that is

$$\ln p_t^* = \alpha_0 + \alpha^e \ln e_t^* + \alpha^p \ln p_t + u_t, \tag{12}$$

where u_t is the usual error term. Equation (12) contains the "error-correction term" u_t that should be stationary according to the Engle-Granger test. However, the current approach differs from the error-correction model of Frankel *et al.* (2012) or Ceglowski (2010) to calculate the ERPT in the sense that u_{t-1} will not enter as an additional variable in a second stage equation such as $\Delta p_t^* = \beta^e \Delta e_t^* + \beta^p \Delta p_t + \lambda u_{t-1} + \epsilon_t$, where λ is an adjustment speed, but rather determines the probability (9) and (10) of being in a low or high-pass regime.

[3]Estimation of the results occurred with Eviews. Chapter 13 of the Eviews manual provides a discussion on switching regressions and the algorithm employed for their estimation. The recursive estimation of the unobserved regime switching regression necessitates the definition of initial probabilities at $t = 0$. Within the present context, these are given by the steady state values implied by the probability transition matrix.

4 Empirical Illustration: Exchange Rate Pass-Through in Switzerland

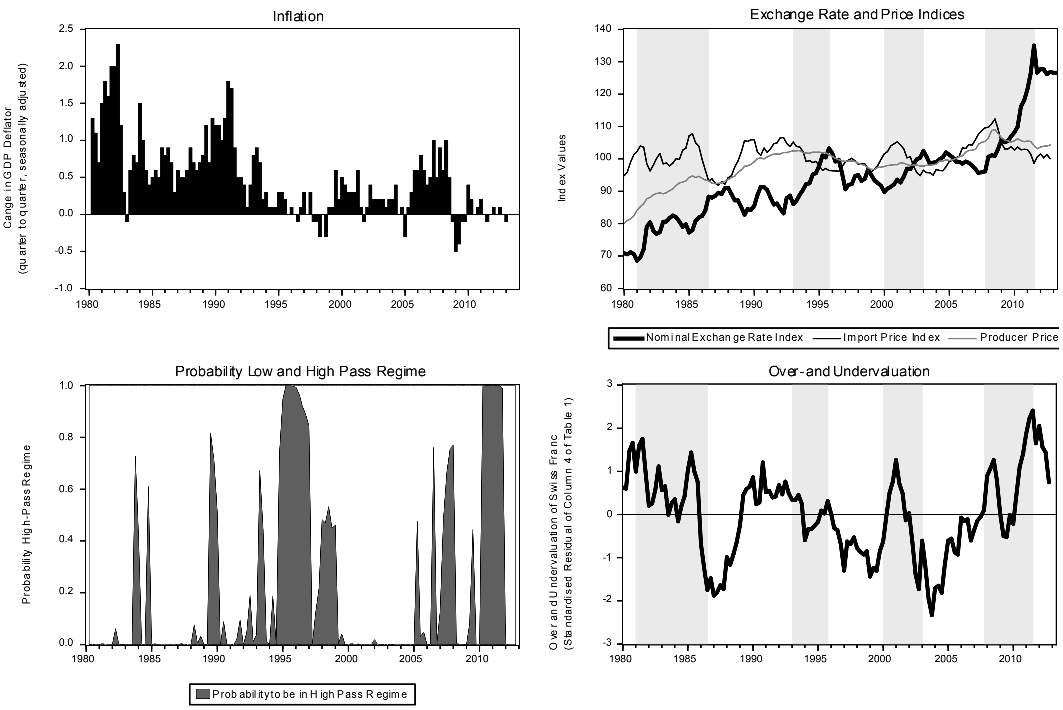
This section provides an example of a regime dependent ERPT by focusing on the case of Switzerland. There are several reasons why Switzerland might lend itself for illustrating the pass-through model of Section 2. Firstly, since the collapse of the Bretton Woods System in the 1970s, the monetary policy conditions have been relatively stable in Switzerland. As depicted in the top left panel of Figure 1, with averages of the quarter-to-quarter change of the GDP Deflator of 0.4 per cent and a maximum value of 2.3 per cent since the beginning of the 1980s, this has lead to a high degree of price stability. This is important since Shintani *et al.* (2013) have shown that, rather than swings in the exchange rate, a time-varying ERPT can also be an artefact of changing levels of inflation. Secondly, as a small economy that is open to international trade, imports account for around 1/3 of Swiss GDP implying that an exchange rate induced shift in the volume of imports could have noticeable price effects. Conversely, Switzerland is too small to affect the world economy, whose condition can therefore be considered as exogenous. Thirdly, though studies dedicated to Switzerland such as Stulz (2007) or Herger (2012) have found a highly incomplete ERPT with typical values between -0.2 and -0.5, the possibility that it could be higher in times of dramatic shifts in the exchange rate has not been considered. Such events can indeed be recurrently observed for the Swiss Franc. In particular, as depicted in the top right panel of Figure 1 by means of a nominal exchange rate index, the Swiss Franc has followed an upward trend against most other currencies. However, due to a safe haven effect, a much steeper appreciation tends to occur in times of international political or financial crises. The most spectacular episode has occurred in the aftermath of the global financial and the Euro crisis when the Swiss Franc appreciated by around 30 per cent within 3 years prompting the Swiss National Bank to set a lower floor of 1.20 against the Euro in September 2011. Owing to this extraordinary effect, in the regression analysis below, a dummy variable will be introduced to account for the possible structural break after the third quarter of 2011. Other episodes where the Swiss Franc appreciated sharply, that are marked by the grey shaded areas, are associated with the bursting of the Dotcom bubble in 2000 and the 9/11 terrorist attacks in 2001, the turbulences in the European Exchange Rate Mechanism in the middle of the 1990s, or the weakening of the US Dollar during the second half of the 1980s.

To connect our results with the widely cited work of Campa and Goldberg (2005), p_t^* is measured by an import price index, e_t by a nominal effective exchange rate index, and p_t by a proxy for local production cost calculated from the product between a producer price index and the ratio between the nominal and real effective exchange rate index. For the case of Switzerland, the top right panel of Figure 1 depicts these variables for the period between the first quarter of 1980 up to the first quarter of 2013. A detailed description of the data and their sources is relegated to the appendix. Campa and Goldberg (2005, p.682) have conducted extensive tests suggesting that these price and exchange rate variables are integrated of order one, that is their (logarithmically) transformed values are non-stationary whilst stationarity arises after a transformation into logarithmic differences, that is Δp_t^*, Δe_t, and Δp_t.[4]

To prepare the field, Table 1 embeds the current sample with some of the previous studies calculating the ERPT. The first column estimates (4) to uncover the short-term pass-through effect. The results concur with the theoretical expectation that an increase in the exchange rate index, which reflects an appreciation of the Swiss Franc, tends to reduce import prices. With an elasticity of -0.48, the magnitude of the coefficient is consistent with the previous findings discussed above. Remarkably, this coefficient estimate is almost identical to the

[4] Conventional tests such as the Augmented Dickey Fuller (ADF) or the Phillips Perron tests on $\ln p_t^*$, $\ln e_t$, and $\ln p_t$ did confirm this finding for the current sample.

Figure 1: Exchange Rates, Prices, and Regime Probabilities



value of $1/2$ that would according to (4) occur under the Cournot-Nash and symmetry assumption ($\nu = 0$ and $\Gamma = 1$). Column 2 extends the specification to the distributed lag model of Campa and Goldberg (2005), where the past four observations, which are supposed to account for the inertia in the ERPT, and ΔGDP enter as additional variables. This hardly changes the instantaneous pass-through effect to -0.46 whilst the long-term impact reflected by the sum of coefficients of the past changes in the exchange rate cumulates to -0.12. Hence, even in the long term, the ERPT remains incomplete. Choudhri and Hakura (2006) have included a lagged dependent variable and a time trend. Again, adopting this specification in column 3 barely affects the value of the instantaneous ERPT.

Ceglowski (2010) and Frankel *et al.* (2012) have employed the error-correction model to disentangle the short and long- term effects of the exchange rate on import prices. Using the single equation approach of error-correction modelling, Column 4 of Table 1 reports the results of the first stage regressing the levels of import prices onto the levels of producer prices and the nominal exchange rate index. The pass-through effect increases to -0.66 but, similar to the long-term effect of the distributed lag model of Column 2, remains statistically different from a compete pass-through. Our measure for the over- or undervaluation of the Swiss Franc, that is according to (12) given by the residual—here reported relative to its standard deviation—of the regression of Column 4, is depicted in the bottom right panel of Figure 1. It is maybe not surprising that periods of overvaluation often overlap with the safe haven episodes marked again by grey shaded areas. Furthermore, with a test statistic of -3.9, the Augmented Dickey Fuller test statistic (ADF) suggests that at the 5 per cent level of rejection, the disequilibrium term is stationary, which permits to include it as an error-correction term in the second stage regression of Column 5. Then again, this does not alter the general picture as regards the value of the short-term pass-through effect.[5] In sum, all specifications of Table 1 give rise to an incomplete pass-through effect in the short

[5] Multivariate time series models yield similar results. In particular, a vector-error-correction-model

Table 1: Different Specifications of the ERPT onto Import Prices in Switzerland

Dependent Variable	Δp_t^*	Δp_t^*	Δp_t^*	$\ln p_t^*$	Δp_t^*
				Error-Correction Model	
				1^{st} Stage	2^{nd} Stage
	(1)	(2)	(3)	(4)	(5)
Δp_{t-1}^*			0.18***		
			(0.06)		
Δe_t	-0.48***	-0.46***	-0.49***		-0.48***
	(0.05)	(0.07)	(0.05)		(0.06)
Δe_{t-1}		-0.11*			
		(0.06)			
Δe_{t-2}		-0.05			
		(0.04)			
Δe_{t-3}		-0.04			
		(0.03)			
Δe_{t-4}		0.08**			
		(0.03)			
Δp_t	0.93***	0.11			0.77***
	(0.15)	(0.18)			(0.13)
Δp_{t-1}		0.57***	0.63***		
		(0.21)	(0.16)		
Δp_{t-2}		0.14			
		(0.28)			
Δp_{t-3}		-0.07			
		(0.21)			
Δp_{t-4}		-0.31***			
		(0.11)			
ΔGDP_t		0.11***			
		(0.04)			
$\ln e_t$				-0.66***	
				(0.11)	
$\ln p_t$				0.63***	
				(0.08)	
u_{t-1}					-0.003**
					(0.001)
T	131	127	131	132	131
R^2	0.54	0.60	0.56	0.63	0.55
AIC	-6.05	-6.03	-6.03	-4.43	-6.05
SIC	-5.90	-5.68	-5.83	-4.30	-5.88

Notes: All specification include seasonal dummy variables and an indicator variable for the period from the third quarter of 2011 onwards to mark the introduction of the lower floor of the Swiss Franc against the Euro. Column (3) includes also a time trend. The data cover the period between 1980 and 2013. Furthermore, T reflects the number of observations, AIC is the Akaike and SIC the value of the Schwarz information criterion. Autocorrelation robust standard errors are reported in parantheses. * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level. the 5% level.

term of about $-1/2$. Furthermore, the information criteria favour the usage of the most parsimonious model of column 1 which will therefore serve as baseline to investigate the possible time varying nature of the (short-term) ERPT.

Table 1 ignores the different regimes of (7) where the short-term ERPT is temporarily higher in times of marked swings in the foreign exchange market. To allow for low and high-pass regimes, Table 2 reports the results of switching regressions. Further to the discussion of Section 3, the various specifications of the switching regressions differ as regards the definition of the transition probability $P[m_t = i]$. In the simplest case, this probability is defined according to (9) and is fixed in terms of depending only on a constant. The corresponding results, reported in column 1, uncover substantial differences in the impact of the exchange rate upon import prices. Specifically, the elasticity almost doubles between the low and high-pass regime from -0.34 to -0.66. Remarkably, the latter value coincides with the ERPT of the second line of (7) when $\nu = 0$ and $\Gamma = 1$. The constant pertaining to the regime switching regression equals -0.07 giving rise to an almost identical probability that an observation is in the high or low-pass regime of, respectively, 0.48 and 0.52. Column 2 employs the Markov-chain of (10) to define the transition probability with respect to a constant. Compared with the results of column 1, the coefficients pertaining to the ERPT change barely. However, in this case, two coefficients arise with respect to the transition between the low and high-pass regimes to represent the matrix with four transition probabilities. The estimated values for this are $(P[m_t = 0|m_{t-1} = 0] = 0.77, P[m_t = 0|m_{t-1} = 1] = 0.23, P[m_t = 1|m_{t-1} = 0] = 0.05, P[m_t = 1|m_{t-1} = 1] = 0.95)$ implying that there is substantial inertia of staying in a given regime.

As discussed in Section 2, if entry competition is responsible for the differential effect of exchange rates on import prices, it is plausible that the transition probability $P[m_t = i|u_{t-1}]$ depends on the degree of over- and undervaluation u_{t-1} of a currency. Using the standardised residual of the bottom right panel of Figure 1, and employing again (9) and (10) to calculate the transition probability, the estimated coefficients are reported, respectively, in columns 3 and 4 of Table 2. Again, a differential effect arises between the low and high-pass regime with similar coefficients to the case with a constant probability. However, u_{t-1} has no significant effect on the regime switching probabilities

Maybe, the specifications of columns 3 and 4 of Table 2 are incomplete in the sense that the transition probability cannot react differently to an over- and undervaluation. Making this distinction could be important for the case of Switzerland since extraordinary effects in the pass-through are mainly associated with the safe haven effect where the Swiss Franc appreciates rapidly. To account for this, for both the simple and Markov switching probability, columns 5 and 6 consider the impact of an undervalued[6] Swiss Franc $u_{t-1} < 0$ and columns 7 and 8 of an overvalued Swiss Franc $u_{t-1} > 0$. Finally, columns 9 and 10 add both effects together in terms of using separate variables for $u_{t-1} < 0$ and $u_{t-1} > 0$ to calculate the regime transition probability. Of note, the short-term pass-through effects remains virtually unchanged across these specifications with an elasticity of around $-1/3$ in the low-pass and around $-2/3$ in the high-pass regime. Based on a t-test with $t = (\beta_0^e - \beta_1^e)/(\sigma_0^e + \sigma_1^e)$, aside from the specifications of columns 4 and 9, the difference in the ERPT is statistically significant at the 10 per cent level of rejection.

The results of Table 2 have been subject to several robustness checks. In particular, as discussed above, the period after the third quarter of 2011 witnessed the extraordinary

(VECM) for $\ln p_t^*, \ln e_t$, and $\ln p_t$ with lag length 1 (which has been chosen by minimising the SIC) did also provide statistical evidence for cointegration.

[6] Of note, with a value of -157.9, the coefficient estimate for δ_{00} in column 6 suggests that the probability $P[m_t = 0|m_{t-1} = 0]$ of remaining in the low-pass regime is essentially equal to 1 when considering the effect of and undervalued currency (where $u_{t-1} < 0$). This concurs with the observation above that dramatic changes in the Swiss Franc exchange rate are mainly associated with appreciations.

Table 2: Regime Switching Regressions of the ERPT onto Import Prices in Switzerland

Transition Probability	Simple (1)	Markov (2)	Simple (3)	Markov (4)	Simple (5)	Markov (6)	Simple (7)	Markov (8)	Simple (9)	Markov (10)
Low Pass Regime										
Δe_t	-0.34***	-0.28***	-0.33***	-0.35***	-0.31***	-0.32***	-0.34***	-0.31***	-0.37***	-0.29***
	(0.05)	(0.07)	(0.05)	(0.06)	(0.03)	(0.03)	(0.05)	(0.08)	(0.06)	(0.06)
Δp_t	1.15***	1.13***	1.16***	1.31***	1.26***	1.14***	1.18***	1.11***	1.05***	1.20***
	(0.19)	(0.10)	(0.20)	(0.18)	(0.23)	(0.19)	(0.22)	(0.24)	(0.18)	(0.18)
High Pass Regime										
Δe_t	-0.66***	-0.63***	-0.65***	-0.56***	-0.64***	-0.69***	-0.63***	-0.68***	-0.52***	-0.64***
	(0.15)	(0.11)	(0.10)	(0.08)	(0.07)	(0.07)	(0.12)	(0.07)	(0.08)	(0.07)
Δp_t	0.45*	0.76***	0.46*	0.31	0.59***	0.46***	0.46**	0.66***	0.64***	0.72***
	(0.24)	(0.13)	(0.25)	(0.20)	(0.19)	(0.17)	(0.22)	(0.17)	(0.20)	(0.15)
Transition Probability										
Constant$(\delta_1;\delta_{00})$	-0.07	1.18*								
	(0.53)	(0.63)								
Constant(δ_{10})		-3.04***								
		(0.66)								
$\upsilon_{t-1}(\delta_1;\delta_{00})$			-0.02	-13.9						
			(0.36)	(26.3)						
$\upsilon_{t-1}(\delta_{10})$				0.41						
				(0.53)						
$\upsilon_{t-1} < 0(\delta_1;\delta_{00})$					-1.80*	-157.9***			3.32*	-3.42***
					(1.08)	(12.63)			(1.88)	(1.34)
$\upsilon_{t-1} < 0(\delta_{10})$						10.25*				-0.56
						(5.13)				(1.78)
$\upsilon_{t-1} > 0(\delta_1;\delta_{00})$							0.20	1.16	-3.66***	9.75*
							(0.72)	(0.78)	(1.12)	(5.46)
$\upsilon_{t-1} > 0(\delta_{10})$								-5.70		-1.67*
								(4.45)		(1.01)
T	131	131	131	131	131	131	131	131	131	131
AIC	-6.04	-6.15	-6.04	-6.08	-6.06	-6.12	-6.04	-6.08	-6.07	-6.16
SIC	-5.77	-5.86	-5.77	-5.80	-5.80	-5.83	-5.77	-5.80	-5.79	-5.83
t-stat(β_r^e)	1.68	1.94	2.13	1.50	3.30	3.70	1.76	2.53	1.07	2.69

Notes: This Table reports the results of regime switching regressions. The simple transition probability is defined by (9) and uses the normalisation $\delta_0 = 0$. The Markov-chain transition probability is defined by (10) and uses the normalisation $\delta_{01} = \delta_{11} = 0$. All specification include seasonal dummy variables and an indicator variable for the period from the third quarter of 2011 onwards to mark the introduction of the lower floor of the Swiss Franc against the Euro. The data cover the period between 1980 and 2013. Furthermore, T reflects the number of observations, AIC is the Akaike and SIC the value of the Schwarz information criterion, and t-stat(β_r^e) is the t-statistic $t = (\beta_0^e - \beta_1^e)/(\sigma_0^e + \sigma_1^e)$ about the difference of the ERPT between the low and high-pass regime. Autocorrelation robust standard errors are reported in parantheses. * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level.

event of the introduction of an exchange rate floor of 1.20 Swiss Francs against the Euro. Likewise, the time after the third quarter of 2008 was marked by the global financial and the Euro crisis with extraordinary circumstances such as interest rates that were close to zero over a long period of time. However, recalculating the results of Table 2 dropping the corresponding observations did not change the essence of the findings. For the sake of brevity, these results are not reported here, but are available on request. Secondly, Shintani *et al.* (2013) have suggested that differences in the ERPT are an artefact of varying levels of inflation. However, the GDP deflator depicted in the top right panel of Figure 1 did not significantly affect the regime switching probability whilst the coefficient estimates for the ERPT remained by and large unchanged.

Based on the comprehensive specification of column 10 of Table 2, which has also lowest value of the AIC, the bottom left panel of Figure 1 depicts the development of the (filtered) probability of being in the low-pass regime. Concurring with the results above, the distinction between the low and high-pass regime follows, by and large, periods during which the Swiss Franc was, respectively, depreciating and appreciating against the major foreign currencies. This result is maybe not surprising since the safe haven effect implies that episodes of a Swiss Franc appreciation arise often amid fundamental shifts compared to episodes of a depreciation that tend to concur with a stable international economy and financial system. In sum, the results of Table 2 provide support for theoretical models where market entries and exits give rise to two regimes with the ERPT being low when fluctuations on the foreign exchange market are modest whilst large values can arise in times of dramatic swings in the foreign exchange market. Econometric models considering only average effect ignore such differences.

5 Summary and Conclusion

The plethora of empirical studies that have been devoted to measuring the exchange rate pass-through (ERPT) find overwhelming evidence that exchange rates impact less than proportional upon import prices. Price inertia arising from menu cost provide one explanation for this incomplete pass-through effect. Al-Abri and Goodwin (2009) and Shintani *et al.* (2013) have recently argued, and confirmed with empirical evidence, that this should introduce nonlinearities in the pass-through with respect to the level of inflation.

Drawing on theoretical work by Baldwin (1988) and Dixit (1989), this paper suggests that nonlinearities in the ERPT can also arise from changes in the market structure. In particular, sufficiently large shifts in foreign exchange rates could trigger market entries or exits which increase the competitive pressure on importing firms to align their prices with currency prices. Hence the hypothesis that "the pass-through coefficient in the phase where the number of foreign firms is constant is quite small [...] whilst its values in the phases with entry or exit are much larger" (Dixit, 1989, p.227). Regime switching regression, where the transition between a low and high-pass regime is a function of the over- and undervaluation of the domestic currency, lend themselves to account for such periodical shifts in the ERPT.

The implication of estimating the ERPT within a regime switching regression is that, even for low inflation countries, temporary increases in the ERPT can occur. This possibility has been confirmed for the case of Switzerland where, despite the generally low level of inflation, spikes in the ERPT seem to occur in times of a dramatic appreciation of the Swiss Franc.

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A Appendix

A.1 Derivation of the Optimal Pricing Rule

Multiplying (2) with E_t/p_t^* and rearranging yields $\pi_t^* E_t/(p_t^*) = [1 - 1/p_t^* e_t]E_t$. Since, $E_t/(p_t^*) = 1/s_t^*$, profits expressed in market share form are given by

$$\pi_t^* = \left[1 - \frac{1}{e_t p_t^*}\right] s_t^*(p_t^*, p_t)E_t.$$

Differentiating this with respect to import prices yields

$$\frac{\partial \pi_t^*}{\partial p_t^*} = \frac{1}{e_t(p_t^*)^2} s_t^* E_t + \left(1 - \frac{1}{e_t p_t^*}\right)\left[\frac{\partial s_t^*}{\partial p_t^*} + \frac{\partial s_t^*}{\partial p_t}\frac{\partial p_t}{\partial p_t^*}\right]E_t = 0.$$

Since (1) defines the import share s_t^*, we have that

$$\left(1 - \frac{1}{e_t p_t^*}\right)\left[-\frac{\gamma^*}{p_t^*} + \frac{\gamma^*\Gamma}{p_t}\frac{\partial p_t}{\partial p_t^*}\right] = -\frac{1}{e_t(p_t^*)^2} s_t^*.$$

Multiplying this with $(p_t^*)^2$ yields

$$(p_t^* - 1/e_t)[-\gamma^* + \gamma^*\nu\Gamma] = -s_t^*/e_t$$

where $\nu = \frac{\partial p_t}{\partial p_t^*}\frac{p_t^*}{p_t}$ is the conjectural elasticity. Solving for the import price yields

$$p_t^* = \left[1 + \frac{s_t^*}{(1-\nu\Gamma)\gamma^*}\right]\frac{1}{e_t}. \tag{13}$$

A.2 Derivation of the Profit Funktion

Inserting (1) into $\widetilde{\pi}_t = (s_t^*/(\gamma^*(1-\nu\Gamma))(1/e_t)$ yields

$$\widetilde{\pi}_t^* = \left[\frac{1}{(1-\nu\Gamma)\gamma^*}\omega_t - \frac{1}{1-\nu\Gamma}\ln p_t^* + \frac{\Gamma}{1-\nu\Gamma}\ln p_t\right](1/e_t).$$

Taking logarithms and using the first order Taylor approximation[7] yields

$$\ln \widetilde{\pi}_t^* \approx \frac{1}{(1-\nu\Gamma)\gamma^*}\omega_t - 1 - \frac{1}{1-\nu\Gamma}\ln p_t^* + \frac{\Gamma}{1-\nu\Gamma} - \ln e_t.$$

A.3 Derivation of High Pass-Through Regime

Inserting $\omega_t = \gamma^* \ln p_t^* - \gamma^*\Gamma \ln p_t + (1-\nu\Gamma)\gamma^* \ln e_t + c$ of the first line of (6) into (4) yields

$$\ln p_t^* \approx \frac{\Gamma}{(2-\Gamma\nu)}\ln p_t - \frac{1-\nu\Gamma}{2-\nu\Gamma}\ln e_t - \frac{1}{1-\Gamma\nu}\ln p_t^* + \frac{\Gamma}{2-\nu\Gamma}\ln p_t - \frac{1-\nu\Gamma}{(2-\nu\Gamma)}\ln e_t + constant$$

Ignoring the constant and solving this for $\ln_t p_t^*$ yields.

$$\Delta p_t^* \approx \big[(2\Gamma)/(3-\nu\Gamma)\big]\ln p_t - \big[(2-2\nu\Gamma)/(3-\nu\Gamma)\big]\ln e_t$$

[7] The first order Taylor approximation of $\ln(x)$ equals $x - 1$ for $0 < x < 2$. This condition is likely to be satisfied for $\ln[\frac{1}{(1-\nu\Gamma)\gamma^*}\omega_t - \frac{1}{1-\nu\Gamma}\ln p_t^* + \frac{\Gamma}{1-\nu\Gamma}\ln p_t]$ since Bergin and Feenstra (2000, p.668) use a parametrisation where γ is 2 whilst, even in small open economies, the share of imports ω_t are much lower than 1 and prices p_t^* and p_t are unlikely to deviate substantially from each other.

B Data Appendix

Table 3: Description of the Data Set

This table describes the variables collected for Switzerland during the 1980 to 2013 period at a quarterly frequency.

Variable	***Unit***	***Description***	***Source***
Dependent Variable:			
Import Price Index (p_t^*)	Index (2005 = 100)	Import price index measured in terms of GDP Deflator (seasonally adjusted values).	Main Economic Indicators (MEI) of OECD. Code: DOBSA.
Independent Variables:			
Producer Price Index (p_t)	Index (2005 = 100)	Following Campa and Goldberg (2005), this is the producer price index (IFS Code: PPI/WPI) multiplied with the ratio between the nominal (IFS Code: NEER) and real effective exchange rate index (IFS Code: REER).	Compiled from International Financial Statistics of IMF.
Exchange Rate Index (e_t)	Index (2005 = 100)	Nominal effective exchange rate index.	International Financial Statistics of IMF. Code: NEER.
GDP	Constant Swiss Francs (2005 = 100)	Real GDP measured by expenditure approach.	Swiss Government, State Secretariat of Economic Affairs.
GDP Deflator	Percent (2005 = 100)	GDP Deflator (quarter on quarter change of seasonally adjusted series). GDP is measured by the expenditure approach.	Swiss Government, State Secretariat of Economic Affairs.

Reviewers Appendix: Additional Results

Table 4: Tests on Stationarity (Compare footnote 7)

Unit-Root Test :	Augmented Dickey Fuller	Phillips Perron		Augmented Dickey Fuller	Phillips Perron
e_t	-2.54	-2.41	Δe_t	-9.51***	-9.52***
p_t^*	-3.85**	-3.35*	Δp_t^*	-8.34***	-8.29***
p_t	-2.79	-2.67	Δp_t	-6.03***	-6.27***

Notes: This table reports the statistics of the Augmented Dickey Fuller (ADF) and the Phillips-Perron test on stationarity. The null hypothesis is that the variables have a unit root. All variables are transformed into logarithms (or logarithmic differences). In the ADF, the lag-length has been selected by minimising the SIC. The tests on the level of the variables contain an intercept and trend. The test on the differences variables contain an intercept (no trend). * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level.

Table 5: Vector Error Correction Model (Compare footnote 8)

	Cointegrating Regression		
p_t^*	-1		
e_t	-0.42***		
	(0.14)		
p_t	0.53***		
	(0.11)		
	Vector Error Correction Model		
	Δp_t^*	Δe_t	Δp_t
Δp_{t-1}^*	-0.20*	0.01	0.08**
	(0.11)	(0.17)	(0.04)
Δe_{t-1}	0.09	-0.19	-0.01
	(0.08)	(0.12)	(0.03)
Δp_{t-1}	0.80***	0.19	0.59***
	(0.21)	(0.32)	(0.07)

Notes: This table reports the results of a VECM. All variables are transformed into logarithms (or logarithmic differences). The lag-length has been selected by minimising the SIC. The error-correction equation contains seasonal dummy variables and an indicator variable for the period from the third quarter of 2011 onwards to mark the introduction of the lower floor of the Swiss Franc against the Euro. The test on the differences variables contain an intercept (no trend). * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level. With a specification with an intercept (no trend) in the cointegrating and vector error correction equation, at the 5% level of rejection, the trace test and maximum eigenvalue tests both find 1 cointegrating relationship.

Table 6: Regime Switching Regression of the ERPT onto Import Prices in Switzerland: Period 1980:Q1 to 2011:Q2

Transition Probability	Simple (1)	Markov (2)	Simple (3)	Markov (4)	Simple (5)	Markov (6)	Simple (7)	Markov (8)	Simple (9)	Markov (10)
Low Pass Regime										
Δe_t	-0.44***	-0.27***	-0.41***	-0.35***	-0.41***	-0.38***	-0.34***	-0.31***	-0.37***	-0.26***
	(0.01)	(0.03)	(0.07)	(0.06)	(0.05)	(0.03)	(0.07)	(0.09)	(0.01)	(0.02)
Δp_t	0.77***	1.27***	1.34***	1.30***	1.23***	1.22***	1.26***	1.10***	1.05***	1.20***
	(0.01)	(0.21)	(0.22)	(0.18)	(0.18)	(0.19)	(0.24)	(0.35)	(0.04)	(0.15)
High Pass Regime										
Δe_t	-0.56***	-0.64***	-0.56***	-0.64***	-0.64***	-0.67***	-0.64***	-0.68***	-0.57***	-0.62***
	(0.05)	(0.06)	(0.08)	(0.08)	(0.10)	(0.07)	(0.08)	(0.08)	(0.05)	(0.06)
Δp_t	0.89***	0.67***	0.48***	0.40	0.25	0.31***	0.53**	0.62*	0.64***	0.70***
	(0.13)	(0.14)	(0.21)	(0.20)	(0.26)	(0.17)	(0.20)	(0.36)	(0.14)	(0.14)
Transition Probability										
Constant($\delta_1;\delta_{00}$)	2.58***	3.39***								
	(0.35)	(0.77)								
Constant(δ_{10})		-2.04***								
		(0.81)								
$\upsilon_{t-1}(\delta_1;\delta_{00})$			-2.85	0.002						
			(2.17)	(26.3)						
$\upsilon_{t-1}(\delta_{10})$				-920.0						
				(0.53)						
$\upsilon_{t-1} < 0(\delta_1;\delta_{00})$					58.93	-0.06			3.19**	-3.55*
					(103.3)	(12.63)			(1.59)	(1.73)
$\upsilon_{t-1} < 0(\delta_{10})$						3.94***				-0.44
						(5.13)				(1.65)
$\upsilon_{t-1} > 0(\delta_1;\delta_{00})$							1.59	3.83	-3.64***	11.52*
							(1.04)	(8.14)	(1.07)	(6.29)
$\upsilon_{t-1} > 0(\delta_{10})$								-0.46		1.55
								(2.99)		(1.18)
T	125	125	125	125	125	125	125	125	125	125
AIC	-6.29	-6.10	-6.04	-6.06	-6.05	-6.05	-6.05	-6.04	-6.26	-6.14
SIC	-6.04	-5.82	-5.80	-5.79	-5.80	-5.77	-5.80	-5.77	-6.00	-5.82

Notes: This Table reports the results of regime switching regressions. The simple transition probability is defined by (9) and uses the normalisation $\delta_0 = 0$. The Markov-chain transition probability is defined by (10) and uses the normalisation $\delta_{01} = \delta_{11} = 0$. All specification include seasonal dummy variables. The data cover the period between 1980 and 2011. Furthermore, T reflects the number of observations, AIC is the Akaike and SIC the value of the Schwarz information criterion. Autocorrelation robust standard errors are reported in parantheses. * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level.

Table 7: Regime Switching Regression of the ERPT onto Import Prices in Switzerland: Period 1980:Q1 to 2008:Q2

Transition Probability	Simple (1)	Markov (2)	Simple (3)	Markov (4)	Simple (5)	Markov (6)	Simple (7)	Markov (8)	Simple (9)	Markov (10)
Low Pass Regime										
Δe_t	-0.44***	-0.31***	-0.41***	-0.30***	-0.34***	-0.38***	-0.37***	-0.37***	-0.37***	-0.33***
	(0.01)	(0.04)	(0.07)	(0.01)	(0.02)	(0.06)	(0.07)	(0.08)	(0.01)	(0.01)
Δp_t	1.09***	0.98***	1.35***	0.95***	1.06***	1.13***	1.28***	1.17***	1.05***	1.09***
	(0.01)	(0.16)	(0.21)	(0.05)	(0.17)	(0.18)	(0.24)	(0.14)	(0.04)	(0.05)
High Pass Regime										
Δe_t	-0.57***	-0.64***	-0.60***	-0.60***	-0.61***	-0.64***	-0.63***	-0.63***	-0.58***	-0.64***
	(0.05)	(0.07)	(0.08)	(0.06)	(0.06)	(0.10)	(0.08)	(0.08)	(0.06)	(0.06)
Δp_t	0.79***	0.64***	0.46***	0.73***	0.65***	0.13	0.45**	0.46**	0.65***	0.67***
	(0.13)	(0.17)	(0.20)	(0.14)	(0.07)	(0.28)	(0.20)	(0.21)	(0.15)	(0.14)
Transition Probability										
Constant$(\delta_1;\delta_{00})$	2.89***	3.37***								
	(0.35)	(1.05)								
Constant(δ_{10})		-2.28***								
		(0.94)								
$\upsilon_{t-1}(\delta_1;\delta_{00})$			-3.45	8.97**						
			(1.94)	(3.61)						
$\upsilon_{t-1}(\delta_{10})$				-0.91						
				(1.28)						
$\upsilon_{t-1} < 0(\delta_1;\delta_{00})$					3.96	-0.90			3.46**	-6.54**
					(2.75)	(0.84)			(1.70)	(3.31)
$\upsilon_{t-1} < 0(\delta_{10})$						85.70				0.08
						(307.1)				(10.45)
$\upsilon_{t-1} > 0(\delta_1;\delta_{00})$							-1.92	1.69	-3.74***	2.39*
							(1.23)	(1.04)	(1.26)	(1.80)
$\upsilon_{t-1} > 0(\delta_{10})$								4.70		1.89
								(10.10)		(2.22)
T	113	113	113	113	113	113	113	113	113	113
AIC	-6.37	-6.10	-6.06	-6.07	-6.08	-6.05	-6.05	-6.03	-6.23	-6.14
SIC	-6.11	-5.81	-5.79	-5.78	-5.82	-5.78	-5.78	-5.74	-5.95	-5.80

Notes: This Table reports the results of regime switching regressions. The simple transition probability is defined by (9) and uses the normalisation $\delta_0 = 0$. The Markov-chain transition probability is defined by (10) and uses the normalisation $\delta_{01} = \delta_{11} = 0$. All specification include seasonal dummy variables. The data cover the period between 1980 and 2008. Furthermore, T reflects the number of observations, AIC is the Akaike and SIC the value of the Schwarz information criterion. Autocorrelation robust standard errors are reported in parantheses. * Significant at the 10% level; ** Significant at the 5% level; *** Significant at the 1% level.